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EARNINGS FORECASTS AND IDIOSYNCRATIC VOLATILITIES

by

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ABSTRACT

Using the Campbell (1991) return decomposition framework, we relate the idiosyncratic volatility of returns to the volatility of changes in expected ROEs for one-, two- and three-year horizon forecasts (i.e., to the volatility of cash-flow news). The upward trend in idiosyncratic volatility documented by Campbell et al. (2001) is associated with a similar increasing trend in the volatilities of cash-flow news for the three forecast horizons. This relationship is persistent after correcting for analysts' forecast biases and is consistent for newly-listed and mature firms and for earnings (non-) announcement periods. Our findings support an informational explanation to the increasing trend in idiosyncratic volatility, and are consistent with the market efficiency hypothesis which implies that stock return volatility is caused by changes in fundamental variables.

Keywords: idiosyncratic volatility, earnings forecasts, market efficiency, cash-flow news, fundamental variables.

JEL Classification: G12; G14; C12; C21; C22; C23.

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1. INTRODUCTION

A fundamental issue in economics, finance and accounting involves the relationship between a firm's earnings and its stock returns. A huge financial accounting literature, which starts with Ball and Brown (1968) and Beaver (1968), investigates the information content of earnings and their impact on realized stock returns and prices. Ball and Brown, among others, demonstrate that accounting earnings capture a portion of the information set that is reflected in security returns.[1] In early tests of market efficiency, Basu (1983) shows that earnings-yield ratios (E/P) help explain the cross-section of average returns on U.S. stocks after controlling for size and market beta.[2] Together with other market anomalies, this evidence led to the three-factor pricing model of Fama and French (1992). Other earnings-related variables that are linked to return predictability include: cash-flow yields and sales growth (Lakonishok et al., 1994), analysts' optimism and forecasts (LaPorta, 1996), dividends to earnings ratios (Lamont, 1998), discretionary accruals (Sloan, 1996; Subramanyam, 1996; Collins and Hribar, 2000a,b), extreme accrual portfolios (Xie, 2001), the probability of changes in earnings (Ou and Penman, 1989a,b), and dispersion in earnings expectations (Diether, Malloy and Scherbina, 2002; Chen, Hong, and Stein, 2002; Park, 2005).

Following the seminal work of Campbell et *al.* (2001), which documents an increase in individual stock's volatility, some researchers examined the empirical relationship between the return variability of individual stocks and changes in fundamentals to investigate whether the upward trend in return volatility can be explained by changes in the uncertainty of the fundamentals. For instance, Vuolteenaho (2000, 2002) uses Campbell's (1991) return decomposition framework to decompose stock returns at the individual-firm level into changes in

[1] Also, see Kormendi and Lipe (1987), Easton and Zmijewski (1989) and Collins and Kothari (1989) for further evidence on earnings response coefficients and the relationship between earnings and stock returns.

[2] Other studies that link return predictability to earnings yields are Shiller (1984) and Fama and French (1988).

cash-flow expectations (cash-flow news) and changes in discount rates (expected-return news). He finds that contrary to aggregate volatility which is mainly driven by expected return news, firm-level stock return variability is highly driven by changes in cash-flow expectations. Wei and Zhang (2006) apply the Vuolteenaho (2000, 2002) methodology and find that the increase in the total volatility of returns on equity helps explain the upward trend in the return volatility of individual stocks. Irvine and Pontiff (2005), Rajgopal and Venkatachalam (2006) and Jiang and Lee (2006) find that measures of cash-flow and/or cash-flow volatility are determinants of idiosyncratic volatility.

The objective of this paper is to examine the relationship between the idiosyncratic volatility of stock returns and the volatility of earnings/cash flows.[3] Our research differs from previous work in four important ways. First, we use the monthly forecasts of analysts instead of reported annual or quarterly earnings to measure volatility. As a result, we do not have to invoke the very tenuous assumption that realized earnings or cash flows on an annual or quarterly basis for the next one to three years are approximately equal to their expected values. Since the forecasts of analysts are available on a monthly basis, this also provides for more timely and synchronous evaluations of the impact of *changes* in earnings expectations on expected returns.[4] Second, our forward-looking focus in on the volatility of *changes* in expected cash flows or cash-flow news as opposed to a backward-looking focus on the volatility of changes in realized earnings. This is based on the belief enunciated by Campbell (1991), among others, that news about changes in the expectations of future cash flows and not their past levels induce subsequent return volatility. Third, our test methodology captures the impact of any forward-looking behavioral biases contained in the expectations of investors (as proxied by the forecasts of analysts) on the idiosyncratic volatility of the returns of individual firms. Fourth, we thoroughly examine the robustness of the relationship

[3] The use of earnings instead of dividends mitigates the impact of dividend smoothing. Although earnings can be manipulated, the extent of earnings manipulation is much lower than for dividends.
[4] Analysts provide more timely forecasts and incorporate information beyond past earnings and financial statements, including market-wide behavior, voluntary disclosures and non-financial information.

between the idiosyncratic volatility of returns and changes in the forecasts of analysts to control for the optimism bias of analysts, firm characteristics (size, leverage and book-to-market) and the heightened volatility associated with earnings announcements.

The use of analysts' forecasts to update expectations about future cash flows is well documented in the literature. Several studies find that analysts' forecast revisions predict future returns (Givoly and Lakonishok, 1980; Stickel, 1991; Liu and Thomas, 2000; Beneish et al., 2001; Gleason and Lee, 2003) and cause a subsequent market response (Griffin, 1976; Givoly and Lokonishok, 1979, 1980; Imhoff and Lobo, 1984; Gleason and Lee, 2003). Research that uses analysts' forecasts to estimate ex ante equity risk premiums documents a strong association between returns and analysts' forecasts (Abarbanell and Bernard, 1992; Frankel and Lee, 1998; Penman and Sougiannis, 1997; Dechow, Hutton and Sloan, 1999; Gebhardt et al., 2001).[5] Roy (1983) finds that the forecasts of both earnings and dividends by analysts are better proxies for market expectations than dividend information alone.

To control for the impact of the forecast bias of analysts,[6] historical measures of forecast error are used to extract the documented forecast bias from the total forecasts of analysts. This allows us to investigate whether the market corrects for the bias when valuing stocks after forecasts are made or updated.[7] Our tests are conducted with(out) a bias correction to assess whether any relationship between idiosyncratic risk and the volatility of updates of earnings forecasts changes after correcting for the bias. Our tests are also conducted with(out) earnings announcement

[5] While there are many advantages when using analysts' forecasts, there is also the problem of the documented bias in the forecasts of analysts (e.g., Gu and Wu, 2000).

[6] Despite their predictive power, many studies report evidence of bias in analysts' forecasts of both short-term (Fried and Givoly, 1982; DeBondt and Thaler, 1990; Abarabanell, 1991; Richardson et al., 1999; Gu and Wu, 2000; Claus and Thomas, 2001) and long-term (LaPorta, 1996; Dechow and Sloan, 1997) earnings forecasts. The evidence on whether this optimism has declined in recent years is conflicting (e.g., Brown, 1997, 1998; Richardson et al, 1999).

[7] See, for example, Michaely and Womack (1999), Dechow, Hutton, and Sloan (1999), Scharfstein and Stein (1990), and LaPorta (1996).

periods to assess whether any increase in idiosyncratic risk is related to the documented evidence of an increasing volatility during such periods (see Landsman and Maydew, 2002).[8]

Our work is consistent with the fundamental view of return variability, which relates return volatility to changes in expectations of either future returns and/or cash-flows. We propose possible explanations related to earnings smoothing, conservatism and/or manipulation.[9] Our results are related to the empirical findings that suggest that the informativeness of financial statements has declined and that the dispersion of analysts' forecasts has increased (Lev and Zarowin, 1999; Rajgopal and Venkatachalam, 2006).[10] A decrease of earnings and financial statements informativeness leads to an increase in cash-flow shocks and thus to an increase in both the volatility of earnings forecasts and the returns of individual firms. However since volatility in analysts' forecasts can also increase with increasing firm risk, our work allows for a test of the informational quality hypothesis against the individual risk hypothesis by comparing how our results vary between small versus large firms and new versus mature firms.[11]

This paper makes at least three contributions to the literature. First, we provide a framework that allows us to explore an explanation for the increasing pattern of idiosyncratic stock return volatility, which is consistent with a fundamental view of return variability and the deterioration of financial reporting. Second, we propose that using earnings forecasts for up to three years as a proxy for earnings expectations provides a more realistic description of how the market adjusts to changes in future cash-flow expectations (Gleason and Lee, 2003) than using annual or quarterly observations of past reported earnings. Third, developing a methodology, which uses both bias

[8] A vast literature that begins with Beaver (1968) reports increased return variability around earnings announcements. Other studies include May (1971), Patell and Wolfson (1984), Lee (1992), Teoh and Wong (1993), Salamon and Stober (1994), and Freeman and Tse (1992).

[9] An increase in cash-flow volatility is related to a deteriorating quality of financial reporting by Rajgopal and Venkatachalam (2006). However, Irvine and Pontiff (2005) relate the increase in cash-flow volatilities to an increasingly competitive environment where firms have less market power and higher fundamental cash-flow shocks.

[10] However, Francis and Schipper (1999) and Landsman and Maydew (2002) find that the information impact of financial statements has increased or stayed constant over time.

[11] Volatility in analysts' forecasts also increases total risk. However, recent work by Campbell et al. (2001) finds no evidence of an increase in total risk.

adjusted and unadjusted earnings forecasts, allows us to examine whether the relationship between idiosyncratic volatility and earnings forecasts is conditional on the bias or whether it is more inherently related to earnings forecasts free from bias.

The remainder of the paper is organized as follows. In the next section, we present the theoretical framework. Section 3 presents the sample and the data. Section 4 presents descriptive statistics and the preliminary analysis. The cross-sectional relationship between earnings volatility and idiosyncratic risk is examined in section 5. Section 6 examines time-series trends in volatility. Section 7 presents robustness checks. Section 8 presents a summary and a discussion of the major findings. Section 9 concludes the paper.

2. THEORETICAL FRAMEWORK

As the efficient market hypothesis implies, a firm's stock return is driven by shocks to expected cash-flows and/or discount rates. Vuolteenaho (2000, 2002) uses the Campbell (1991) return-decomposition framework and an accounting-based approach known as the clean surplus relationship to develop such implications for stocks of individual firms. This leads to the following decomposition, where return-on-equity is used instead of dividends to represent the relationship between returns and cash-flow changes:[12]

$$r_t - E_{t-1}r_t = \Delta E_t \sum_{j=0}^{\infty} \rho^j \left(e_{t+j} - f_{t+j} \right) - \Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} + \kappa_t \quad (1)$$

where ΔE_t denotes the change in expectations from *t-1* to *t*; $e_t = \log\left(1 + X_t / B_{t-1}\right)$ is the return-on-equity from *t-1* to *t*; $r_t = \log\left(1 + R_t + F_t\right) - f_t$; $f_t = \log(1 + F_t)$; R_t is the simple excess stock return; F_t is the risk-free interest rate; X_t is earnings from *t-1* to *t*; ρ is a constant slightly less than one; B_{t-1} is the book-value of equity at *t-1*; and κ_t is an approximation error. Using (1), the

[12] Liu, Thomas and Nissim (2002, 2006) find that cash-flow valuations are dominated by earnings, and that earnings-based valuations are closer to traded prices than cash-flow-based valuations.

unexpected return variance is decomposed into the variance of cash-flow news and the variance of expected return news to obtain:

$$\mathrm{var}\left(r_t - E_{t-1}r_t\right) = \mathrm{var}\left(N_{r,t}\right) + \mathrm{var}\left(N_{cf,t}\right) - 2\,\mathrm{cov}\left(N_{r,t}, N_{cf,t}\right), \text{ and} \tag{2}$$

$$N_{cf,t} \equiv \Delta E_t \sum_{j=0}^{\infty} \rho^j \left(e_{t+j} - f_{t+j}\right) + \kappa_t, \text{ and } N_{r,t} \equiv \Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j}. \tag{3}$$

Using a VAR approach, Vuolteenaho finds that firm level returns are mainly driven by cash-flow news rather than expected return news. Wei and Zhang (2006) use the Vuolteenaho decomposition and focus on the relationship between conditional return volatility and the conditional variance of cash-flow news. They find that stock return volatility is positively related to the volatility of the return on equity.

Our work also aims at studying the relationship between idiosyncratic volatility and changes in expected cash-flows. Similar to Wei and Zhang, we focus on the (conditional) variance of cash-flows when studying return volatility. Specifically, we use:

$$\mathrm{var}\left(r_t - E_{t-1}r_t\right) = \mathrm{var}\left[\Delta E_t \sum_{j=1}^{\infty} \rho^j \left(e_{t+j}\right)\right] + \zeta_t \tag{4}$$

where ζ_t contains the variances in the expected return news, the error term κ_t, the risk-free rate and all the covariance terms.[13] However, unlike Wei and Zhang (2006) who use the time-series of quarterly earnings forecasts to measure the conditional volatility of return on equity, our approach focuses on revisions of analyst's forecasts of earnings and their impact on idiosyncratic volatility. This is based on the belief that the revision and hence the news about future cash flows and not their level causes returns to change (Campbell, 1991). This allows us to include revisions in forecasts of multi-year-ahead earnings, which should explain more variation in stock returns (Liu

[13] Vuolteenaho finds that the error term κ_t is small and that the covariance between cash-flow news and discount rate news is also small. Campbell, Lo and Mackinlay (1997) assume that news about future dividends are uncorrelated with news about future returns. They assert that this assumption could be true if, for example, expected returns are determined by the volatility of the dividend growth process and dividend volatility is driven by a GARCH model so that shocks to volatility and the level of dividends are uncorrelated.

and Thomas, 2000). The empirical model used in our tests includes forecasts for the next three years. This leads to the following approximate relationship:

$$\begin{aligned}\operatorname{var}\left(r_t - E_{t-1} r_t\right) &= \operatorname{var}\left[\Delta E_t \rho e_{t+1}\right] + \operatorname{var}\left[\Delta E_t \rho^2 e_{t+2}\right] + \operatorname{var}\left[\Delta E_t \rho^3 e_{t+3}\right] \\ &+2\operatorname{cov}\left[\Delta E_t \rho e_{t+1}, \Delta E_t \rho^2 e_{t+2}\right] + 2\operatorname{cov}\left[\Delta E_t \rho e_{t+1}, \Delta E_t \rho^3 e_{t+3}\right] \\ &+2\operatorname{cov}\left[\Delta E_t \rho^2 e_{t+2}, \Delta E_t \rho^3 e_{t+3}\right] + \psi_t + \zeta_t\end{aligned} \tag{5}$$

where ψ_t encompasses variance terms of higher-period forecasts. We assume that ψ_t and ζ_t are uncorrelated.

Equation (5) is then used to assess how idiosyncratic return volatility is related to revisions in expected cash flows because revisions of analysts' forecasts provide a more comprehensive, frequent and timely source of information for investors to use when updating their expectations about stock returns. The revision of analysts' forecasts also plays an important role in the diffusion of information about future corporate earnings and unlike earnings announcements, which occur at a quarterly basis, the revisions of analysts' forecasts are made on a monthly or higher frequency basis.

Many studies ignore the biases in the forecasts of analysts by, in effect, implicitly assuming that investors naively rely on the forecasts of analysts (e.g., LaPorta, 1996; Dechow and Sloan, 1997; Jiambalvo, Myers and Peecher, 1998). In turn, this implies that investors are not fully rational towards the bias or that its magnitude is completely unpredictable. Other studies report evidence that investors make adjustments for predictable bias (e.g., Freeman and Tse, 1992; Dugar and Nathan, 1995, Han, Manry and Shaw, 2001).

Herein, we assume that investors use a simple approach to adjust analysts' forecasts for predictable bias. Specifically, we assume that investors extrapolate average past biases and use them as an indicator of future biases. To reduce the impact of forecast error in bias estimation, the last forecast made before an earnings announcement is used as our proxy of the least biased analysts' forecast.[14] This procedure has two advantages. First, it helps in differentiating the forecast bias from the forecast error term, since the latter is induced by innovations to expected earnings and should be independent from the forecast bias. Second, using the last period forecast instead of reported earnings, as a proxy of the least biased forecast, mitigates the impact of management manipulation which affects reported earnings and could only distort the value of the bias if considered. Since analysts make forecasts of the fundamental value of earnings, which is not always equal to the reported value, the forecast bias is defined as:

$$bias_{i,m,t} = fe_{i,l,t-1} - fe_{i,m,t-1} \tag{6}$$

$$perbias_{i,m,t} = \left(fe_{i,l,t-1} - fe_{i,m,t-1} \right) / fe_{i,m,t-1} \tag{7}$$

$$adfe_{i,m,t} = fe_{i,m,t} \left(1 - avg\_perbias_{i,m,t} \right) \tag{8}$$

$$avg\_perbias_{i,m,t} = \sum_{k=t-3}^{t-1} perbias_{i,m,k} \tag{9}$$

where $bias_{i,m,t}$ is the forecast bias for month *m* of year *t* based on observed forecasts at year *t-1*; $perbias_{i,m,t}$ is the percentage bias forecast as measured with respect to the forecast made during month *m* of year *t-1*; $adfe_{i,m,t}$ is the adjusted forecast for month *m* of year *t* that controls for the bias of analysts; $fe_{i,l,t-1}$ is the last (*l*) analysts' forecast made before earnings of firm *i* at the end of fiscal year *t-1* are released; $fe_{i,m,t-1}$ is analysts' forecast for month *m* for earnings of firm *i* at the end of fiscal year *t-1*; $adfe_{i,m,t}$ is our measure of the adjusted forecasts that is subsequently used in the empirical work; and $avg\_perbias_{i,m,t}$ is the average forecast error for month *m* over the past three years.[15] Forecast biases for years 2 and 3 are constructed in a similar fashion.[16]

[14] We are assuming that forecasts prior to the last forecast contain low forecast error and high forecast bias (if any).

[15] The averaging procedure used when estimating the bias could underestimate the bias since positive and negative biases of similar magnitudes average out.

[16] This method leads to the elimination of the one-year ahead bias only. Two- and three-year-ahead biases are not eliminated.

3. SAMPLE AND DATA

The data used herein is collected from three sources: daily returns from CRSP, book-value of equity from Compustat, and monthly consensus analysts' forecasts from I/B/E/S. The sample period extends from 1976 (the beginning of I/B/E/S coverage) to 2003, when we use a two-period forecast horizon only, and is restricted to 1982 to 2003 when we include the three-year forecast horizon, since these forecasts are only available starting from 1982. To maintain accounting consistency between firms, only stocks with calendar fiscal year-ends are considered.[17] A stock is included in the sample if it has positive book-value of equity, return data in a particular month and a series of monthly earnings forecasts for one-, two- and three-years. When three-year-hence earnings' forecasts are not available, we use the long-term growth forecasts, whenever available, to construct the series of monthly three-year forecasts. For each month, both adjusted and unadjusted earnings forecasts are computed where adjusted forecasts are based on the bias correction for the past three years, as described in equations (6)-(9) above.

Changes in (un)adjusted median monthly earnings forecasts are used to calculate the sample variance of changes in expected returns on equity (EROE), which is given by the monthly EPS forecast multiplied by the number of common shares used to calculate EPS (data item 54 in Compustat) and divided by the book value of equity. The book value of equity is the total common equity (data item 60 in Compustat) at the beginning of the year. The last change in monthly forecasts is computed with respect to reported earnings, which are computed using Net Income (Compustat data item 172). Dates of reported earnings are deduced from IBES, and they correspond to the months when forecasts shift horizons. This usually occurs three months after the end of the fiscal year. To be consistent with the idiosyncratic volatility literature, which usually computes excess return values, all ROE values are computed in terms of excess values relative to

[17] This restriction is also imposed since earnings forecast adjustment models, which depend on the distribution of prior earnings changes, would be easier to estimate in the cross-section when all firms are aligned in calendar time.

the monthly risk-free rate. To reduce the impact of outliers and to avoid spurious inferences due to extreme values, the series of expected ROE (EROE), adjusted expected ROE (AEROE) and reported ROE are winsorized at -100% and +100%. The sample variance of the changes in EROE (AEROE) is computed using a trailing 12-month window for each month and each forecast horizon. Due to this requirement, our variance measures start from 1977 for the one- and two-year forecasts and from 1983 for the three-year forecasts.

4. DESCRIPTIVE STATISTICS

Table 1 reports descriptive statistics for the volatility and covariance of unadjusted monthly changes in EROE. In order to assess how the volatility changes over time, the entire sample period is divided into several five-year subperiods. For each subperiod, the time-series average of the volatility and the covariance of changes in EROE across all firms are computed for all months in the subperiod. Panel A reports equal- and value-weighted volatilities of changes in EROE for one-, two- and three-year horizons for all CRSP firms that satisfy our data requirements. The average number of firms for which one-year IBES forecasts exist is about 953 during the subperiod 1977-1981 but increases to reach 3476 during the 1996-2000 subperiod. The latter figure represents more than 90% of the available CRSP firms with December fiscal year-ends for that subperiod. The number of firms for which two-year (three-year) IBES forecasts are available is comparable (more modest) and shows a similar pattern of growth. The average number of firms for which we have one-, two- or three-year forecasts over the whole sample period are 2081, 1994 and 932, respectively.

[Please place Table 1 about here]

The statistics reported in panel A of Table 1 show that the variance of changes in EROE exhibits an increasing pattern for the one- and two-year (not three-year) forecast horizons. The mean variance of changes in one-year expected ROE more than quadruples from 1977-1980 to 2001-2003. The same phenomenon is observed for the two-year horizon forecasts. These results are consistent with the findings of Wei and Zhang (2006) and Irvine and Pontiff (2005) who also document an increase in the volatility of ROE and cash-flows, respectively. Many explanations could be offered for such an increase. For instance, this could indicate an increase in informational uncertainty in the investment environment, or the inclusion of smaller firms which were not as well represented in the earlier subperiods. The average variance of changes in EROE for the one-year horizon is usually lower than that for the two- and three-year horizons. The only exception is the 1977-1980 subperiod, where the volatility of changes in one-year EROE was almost double the volatility of changes in two-year EROE. The consistent increase in the volatility of the three EROE variables starts to revert in the last subperiod, 2001-2003, which is consistent with the findings of Brandt, Brav and Graham (2005) who document that idiosyncratic volatility measures have declined from 2000 to 2004 to levels commonly observed in the 1970s and early 1980s. The equal-weighted averages of changes in EROE volatilities are higher than their value-weighted counterparts, indicating that small firms have more volatile changes in their earnings and ROE forecasts.

Panel B of Table 1 reports descriptive statistics for the covariances of changes in EROE for the one-, two- and three-year horizons. All the mean covariances are positive, which indicates that updates in forecasts usually take the same direction for different horizon forecasts. This provides some support for the evidence of earnings persistence. Unlike the three-year related covariances (COV13 and COV23), the covariance between the changes in one- and two-year forecasts (COV12) shows an increasing trend. This suggests that the impact of news has increasingly a more correlated impact for both the one- and two-year horizons, and a neutral impact for the three-year horizon.

Panel C of Table 1 reports the correlation matrix between volatilities and the covariances for the different horizons. One correlation value exceeds 0.8 and three exceed 0.7. This indicates that our results should be interpreted with some caution, and that we should examine them based on the overall (not individual) impact of different horizon forecasts on idiosyncratic volatility.

The patterns presented in Table 1 are also illustrated graphically. Panel A of Figure 1 presents the time-series plots of the average monthly equal-weighted volatilities of unadjusted changes in ROE expectations for one-, two- and three-year forecasts (i.e., $V\Delta EROE1_{ew}$, $V\Delta EROE2_{ew}$, and $V\Delta EROE3_{ew}$, respectively). As noted earlier, the plots show a clear increasing trend for the one- and two-year forecast series. The three-year forecast series is characterized by higher swings around a somewhat increasing trend, and some seasonality that might reflect the surprise effect around the earnings announcement dates. All the series of volatilities exhibit a decrease in volatility starting from 2002. This trend is also shared by the idiosyncratic volatility as is shown subsequently in Figure 3.

[Please place Figure 1 about here]

Panel B of Figure 1 presents the time-series plots of the value-weighted volatilities of the unadjusted changes in ROE expectations for one-, two- and three-year forecasts ($V\Delta EROE1_{vw}$, $V\Delta EROE2_{vw}$, and $V\Delta EROE3_{vw}$, respectively). All the series show lower values of volatility compared to their equal-weighted counterparts, which suggests that changes in forecasts are more volatile for small versus large firms. Similar to the equal-weighted volatility measures, the value-weighted volatilities for one- and two-year forecasts have an increasing trend with some evidence of seasonality for the two- (and three-) year forecasts. As is the case for the equal-weighted counterparts, seasonality and large swings around a somewhat increasing trend up to 2001 are also clearly observed in the three-year forecasts of the value-weighted volatility, $V\Delta EROE3_{vw}$.

Table 2 reports descriptive statistics on the volatility of the bias-adjusted changes in EROE expectations, AEROE. The mean equal- and value-weighted volatilities of bias-adjusted changes in ROE expectations for all horizon forecasts and sample periods are higher than their unadjusted counterparts (Panel A of Table 2). As expected, this indicates that biased forecasts are smoother and less volatile than their corrected counterparts. The increasing trends observed in the unadjusted volatilities are also present in the bias-adjusted volatility measures. Based on Panel C of Table 2, the correlations between the bias-adjusted volatilities and covariance variables are lower on average than their unadjusted counterparts. Based on Panel D of Table 2, both the unadjusted and bias-adjusted changes in ROE expectations have similar covariances.

[Please place Table 2 about here]

As depicted in Panels A and B of Figure 2, both the equal- and value-weighted bias-adjusted volatilities have an increasing trend for one-, two- and three-year forecast horizons. Since the trend is modest for the value-weighted volatility measures, this indicates that an important part of the trend is related to smaller firms.

[Please place Figure 2 about here]

We measure idiosyncratic volatilities for individual stocks using the single-factor CAPM market model as well as the Carhart (1997) model, which includes the three Fama and French (1993) factors and a momentum factor. In every month, daily excess returns of individual stocks are regressed as follows:

$$r_{i,t} = \alpha + \beta_{i,1} \cdot MKT_t + \varepsilon_{i,t}^{1F} \tag{10}$$

$$r_{i,t} = \alpha + \beta_{i,1} \cdot MKT_t + \beta_{i,2} \cdot SMB_t + \beta_{i,3} \cdot HML_t + \beta_{i,4} \cdot MOML_t + \varepsilon_{i,t}^{4F} \tag{11}$$

where $MKT_t = R_{m,t} - R_{f,t}$, *SMB* is the difference between the returns on portfolios of small and large stocks, *HML* is the difference of the returns on portfolios of high and low book-to-market stocks, and *MOM* is the difference in returns between portfolios of high and low prior return stocks, as downloaded from the website of Kenneth R. French. The time-series regressions are run for each stock for each month, and the idiosyncratic volatility of a stock is computed as the standard deviation of the regression residuals. Similar to Fu (2005), we require a minimum of 15 days per month for which daily return and non-zero trading volumes are reported on CRSP. Each variance of daily return residuals is multiplied by the number of trading days for each month in order to get an estimate of the stock's variance for that month.

In order to assess how volatility changes over time, the entire sample period is divided into five-year subperiods. The time-series averages of idiosyncratic volatilities across all firms are computed for all months in each subperiod. Based on Panel A of Table 3, the number of firms ranges from 1566 in 1977-1980 to 6346 in 2001-2005. Consistent with previous work, the means of the monthly equal- and value-weighted idiosyncratic volatilities for the all-firm sample have an increasing trend over the five-year subperiods from 1977 to 2000. The values range from 1% to 4.7% for IV^{1f}_{ew} and from 0.3% to 1.4% for IV^{1f}_{vw}. This trend reverses in the last subperiod (2001-2005). The value-weighted volatility mean values are systematically lower than their equal-weighted counterparts indicating that larger firms have, on average, lower idiosyncratic volatilities. These results are consistent with the Campbell *et al.* (2001) findings that the increase in idiosyncratic volatility prior to the year 2000 is not caused only by small firms. Both measures of volatility are positively skewed with positive kurtosis. Similar to previous studies, results based on the 3- and 4-factor models are quite comparable.

[Please place Table 3 about here]

Descriptive statistics for the idiosyncratic volatilities of CRSP firms that satisfy our data requirements for IBES forecasts are reported in Panel B of Table 3. The number of such firms ranges from 712 during 1977-1980 to 2971 during 2001-2005 with an average number of 1914 over the whole sample period. For this sample, we find that the descriptive statistics for both the equal- and value-weighted idiosyncratic return volatilities, $IV^{1f}_{ew,ibes}$ and $IV^{1f}_{vw,ibes}$, are comparable to the descriptive statistics for all CRSP firms (i.e., for IV^{1f}_{ew} and IV^{1f}_{vw}, respectively).

The time-series patterns presented in Table 3 are illustrated graphically in Figure 3 for $IV^{1f}_{ew,ibes}$ and $IV^{1f}_{vw,ibes}$ based only on the single-market factor since they are similar to those from the 4-factor model and for all CRSP firms. Both series show an increasing trend in volatility, with a higher slope for the equal-weighted measure. The increase is moderate prior to 1987 and becomes more pronounced during the 1990s for both series. As noted previously, the trend reverses after 2001 indicating a sharp decrease in idiosyncratic volatilities.

[Please place Figure 3 about here]

5. CROSS-SECTIONAL REGRESSION RESULTS

The descriptive statistics and preliminary analysis discussed in the previous section show that the idiosyncratic volatilities of stock returns are increasing over time, and that the volatilities of bias (un)adjusted changes in ROE expectations for one- and two-year forecasts (and to a lesser extent for their three-year counterparts) are also increasing over time. The cross-sectional analysis conducted in this section allows us to better understand the dynamics of any relationship between the volatility of changes in expected ROE and idiosyncratic volatility, and to make sure that any time-series relationship is not spurious.

At the end of each month, the return idiosyncratic volatility is regressed against the bias unadjusted changes in expected ROE for one-, two- and three-year horizons, as described previously in equation (7). This leads to the following empirical formulation:[18]

$$\begin{aligned} IV_{i,t}^{4f} &= \beta_0 + \beta_1 * \rho^{\tau} * V\Delta EROE1_{i,t} + \beta_2 * \rho^{\tau+1} * V\Delta EROE2_{i,t} \\ &+ \beta_3 * \rho^{\tau+2} * V\Delta EROE3_{i,t} + \beta_4 * \rho^{2\tau+1} * COV12_{i,t} \\ &+ \beta_5 * \rho^{2\tau+2} * COV13_{i,t} + \beta_6 * \rho^{2\tau+3} * COV23_{i,t} + \varepsilon_{i,t} \end{aligned} \tag{12}$$

To reduce the impact of cross-sectional correlations, we use the Fama and MacBeth procedure and average the slopes of the monthly regressions and their time-series standards errors to draw inferences. The Newey and West (1987) procedure is used to correct the t-statistics for autocorrelation and conditional heteroscedasticity in the return volatilities and the changes in EROE volatilities for the different forecast horizons. To capture the impact of discounting, we assume a constant value of ρ, ρ=0.95.[19] The impact of any within-year discounting for time value is captured by τ , or the number of months before the next fiscal year-end divided by 12. This allows us to capture the changes in the sensitivities of the return volatilities to the volatilities of

[18] The idiosyncratic volatility of returns could also be related to changes in the dispersion of ROE based on changes in the dispersions of the forecasts of analysts. Prior studies argue for the use of the dispersion of the forecasts of analysts as a proxy for information asymmetry and the quality of firms. disclosures (e.g., Krishnaswami and Subramaniam, 1998). However, this is left to future analysis.

[19] This is the standard assumption in the literature (e.g., Campbell, 1991). Although discount rate estimation introduces noise into the parameters estimation, we believe that its impact is small and would not alter our major conclusions.

forecast changes as the EPS reporting date approaches.[20] A similar procedure is used to examine the relationship between idiosyncratic volatilities and the volatilities of one-, two- and three-year bias-adjusted changes in expected ROE.

Regression results related to the bias-unadjusted changes in expected ROE based on the four-factor idiosyncratic volatility model using two different specifications are reported in Table 4.[21] The restricted model specification assumes that the market only considers changes in the closest two forecasts to assess the impact of changes in all future ROE expectations on return volatility. The regressions for this specification are first conducted over the whole sample period (1977-2003), and then over successive five-year subperiods to assess how the relationships change over time. Based on Panel A of Table 4, the relationship between monthly return idiosyncratic volatilities and the volatilities of monthly changes in forecasts of both one- and two-year ROE is significant and positive over the whole time period and the subperiods. The relationship is significant and negative for the covariance term between the two series over the whole period and subperiods (but only significant for the 1977-1980 and 1981-1985 subperiods). A positive covariance in forecast changes implies market confidence in its forecasts and can be perceived as a signal which reduces return volatility. The R^2 values range from 3% to 7% for the subperiods, and the R^2 is 5% for the whole time period.

[Please place Table 4 about here]

The regression results that also include the three-year forecasts are summarized in Panel B of Table 4. Since IBES starts recording analysts' LTG and three-year forecasts in 1982, these results are for the restricted time period of 1985-2003. Idiosyncratic volatility continues to be positively related to the volatilities of changes of monthly forecasts of ROE expectations for this less restricted specification for the whole (but shorter) time period and for all but one subperiod. For the 1986-1990 subperiod, the volatility of monthly changes in one-year ROE expectations and the

[20] However, results and inferences do not differ substantially if we do not correct for this change in the discounting effect.
[21] Results based on the single-factor model are similar.

volatility of monthly changes in three-year ROE expectations are both non-significant. The relationship is significant (and negative) with the covariance between changes in one- and two-year forecasts and with the covariance between changes in two- and three-year forecasts, and not significant with the covariance between changes in one- and three-year forecasts. The covariance terms are not consistently significant across all subperiods. The only exception is the covariance between changes in ROE expectations of one- and three-year forecasts, which is consistently negative and significant across all subperiods. These results show that monthly changes in three-year ROE expectations have an impact on the volatility of stocks. The negative covariances can be caused by time-varying reversals of the accrual errors that are incorporated in ROE expectations. The overall model improves in terms of fit when changes in three-year forecasts are also included. The overall R^2 goes from 0.05 to 0.11 and ranges from 0.05 to 0.11 for the different subperiods.

The regression results when changes in ROE forecasts are corrected for analysts' bias using the method as described in equations (6) to (9) are reported in Table 5. The overall results are similar to the ones reported earlier in Panels A and B of Table 4. They confirm that there is a positive relationship between monthly idiosyncratic volatility and volatility of changes in monthly forecasts of one-, two- and three-year forecasts of ROE. However, the impact of the covariance terms is lower in Panel B of Table 5 for this less restricted specification. The results across the different subperiods are also comparable although more instances occur where the independent variables are less significant and where the coefficient values are somewhat lower. This seems to indicate that the market does not easily detect the bias in earnings forecasts and reacts more to the bias-unadjusted forecasts, which leads to an argument in support of the deteriorating quality of financial reporting as an explanatory variable of the increase in volatility. However, the overall model significance does increase when bias-adjusted variables are used, which indicates that the bias-corrected model has the higher explanatory power.

[Please place Table 5 about here]

Overall, the cross-sectional results indicate a positive and significant relationship between idiosyncratic volatility of stock return and volatility of changes in monthly forecasts of expected ROE for one-, two-, and three-year horizons, or which could be described as the volatility in cash-flow news.[22] This relationship is detected through the whole sample period but is also consistent throughout the different five-year sub-periods, indicating that the association is consistent through time.

6. TREND AND TIME-SERIES RESULTS

We now conduct a trend analysis to ensure that the cross-sectional relationship is consistent through time and a time-series association exists between the variables. The following cross-sectional time-series regression which controls for time-trend in volatility is conducted, where regressions are estimated using the GMM method and the White standard error correction:[23]

$$\begin{aligned} IV_{i,t}^{4f} &= \delta_0 + \delta_1 * time + \delta_2 * \rho^{\tau} * V\Delta EROE1_{i,t} + \delta_3 * \rho^{\tau+1} * V\Delta EROE2_{i,t} \\ &+ \delta_4 * \rho^{\tau+2} * V\Delta EROE3_{i,t} + \delta_5 * \rho^{2\tau+1} * COV12_{i,t} \\ &+ \delta_6 * \rho^{2\tau+2} * COV13_{i,t} + \delta_7 * \rho^{2\tau+3} * COV23_{i,t} + \xi_{i,t} \end{aligned} \tag{13}$$

Panel A of Table 6 reports the regression results when using the bias unadjusted changes in expected ROE for one-, two- and three-year horizons. The first regression uses the 1977-2003 time period and only one- and two-year horizon changes in expected ROE plus the trend. The trend slope coefficient is positive and significant confirming the existence of an increasing trend in stock return idiosyncratic volatilities. However, this does not subsume the significance of the ROE volatility-related variables. The coefficients for the one- and two-year horizon changes in expected ROE volatilities are both positive and significant, confirming the robustness across time of the cross-sectional regression results. The coefficient on the covariance variable is also positive

[22] The two terms are not exactly equivalent since the cash flows are adjusted by the value of equity in the former case which would affect the volatility of cash flows by the changes in equity.
[23] We use firms' fixed effects to control for omitted variables which could be correlated with the explanatory variables.

but not significant. The coefficients of the variables and their significance remain quite similar when the sample period is restricted due to data availability to 1985-2003.

[Please place Table 6 about here]

When the three-year horizon volatility of changes in ROE is added, the one-year horizon volatilities coefficient and its significance are slightly reduced while the three-year horizon volatility coefficient which has a positive sign is not significant. The two phenomena could be due to the correlation between the two explanatory variables, since the significance of the overall model slightly increases when we add the three-year explanatory variable. None of the cross-sectional correlation variables is significant. The significance of the overall time-series model ranges from 20% to 22%, which is much higher than the cross-sectional model, most likely because it is capturing both cross-sectional and time-series effects.

Panel B of Table 6 reports the regression results when using the bias-adjusted changes in expected ROE for one-, two- and three-year horizons. Compared to Panel A results, the sign of the variables and the significance of their coefficients are maintained except for the covariance variables which become negative but remain not significant. The coefficients of all but the time variable are consistently lower (as at the cross-sectional level). This could be due to the market fixating on non-bias versus bias-corrected forecasts.[24] However, the overall model significance is higher when we control for the bias which is consistent with the view that the market understands and corrects somewhat for the bias. The overall time-series results indicate that even after controlling for the trend effect in idiosyncratic volatility, variability in changes in expected ROE plays a role in explaining the upward trend in idiosyncratic volatility and accounts for a big part of the increase in idiosyncratic volatility.

We also conduct a series of unit root tests, which are reported in Table 7, to verify that all the variables studied are not integrated. To better capture the individual movements of firms, we use

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[24] We do not exclude the possibility that our naïve method of bias correction is responsible for the lower coefficients.

panel-based unit root tests since they have higher power and provide more information than unit root tests based on aggregate time series. This procedure has the advantage of testing for unit root or persistence in individual cross-sectional time series unlike the unit-root tests usually applied on average/aggregate time-series without testing for individual persistence and specification. Although the latter tests provide good results when examining aggregate persistence, they are not appropriate for detecting individual persistence.

The following three types of panel unit root tests are used based on their performance and efficiency:[25] Levin, Lin and Chu (2002), Breitung (2000) and Im, Pesaran, and Shin (2003). The first two assume a common autoregressive structure for all the series, whereas the third test allows for different autoregressive coefficients in each series. Since our previous results show the existence of a trend, we only conduct the unit root tests with a trend, as we would expect the non-trend tests to provide support for a unit root. To control for autocorrelation in volatilities, we set the number of lags to 4. Table 7 reports unit root results on stock return idiosyncratic volatility as well as the volatility of monthly changes in ROE expectations for one-, two-, and three-year horizons for bias-(un)adjusted forecasts. The three sets of test results show evidence of the absence of a unit root for the stock return idiosyncratic volatility. The results for the volatility of changes in ROE expectations are usually supportive of a unit root. Mixed evidence for a unit root exist for bias-unadjusted volatility for the three-year horizon and bias-adjusted volatility for the one-year horizon.

[Please place Table 7 about here]

7. ROBUSTNESS CHECKS AND OTHER TESTS

Recent studies argue that newly listed firms could be responsible for the increase in idiosyncratic volatility (Wei and Zhang, 2006; Irvine and Pontiff, 2005; Fink, Fink, Grullon and Weston, 2005; Fama and French, 2004). If such is the case, we expect that the different volatility

[25] See, for instance, Baltagi (2005), and Hlouskova and Wagner (2005).

measures would be more stable after controlling for newly-listed firms. The importance of newly-listed firms in the increasing trend of idiosyncratic volatility is investigated by introducing a dummy variable into the following time-series cross-sectional regression:[26]

$$\begin{aligned}
IV_{i,t}^{4f} &= \delta_0 + \delta_1 * time + \delta_2 * \rho^{\tau} * V\Delta EROE1_{i,t} + \delta_3 * \rho^{\tau+1} * V\Delta EROE2_{i,t} \\
&+ \delta_4 * \rho^{\tau+2} * V\Delta EROE3_{i,t} + \delta_5 * \rho^{2\tau+1} * COV12_{i,t} + \delta_6 * \rho^{2\tau+2} * COV13_{i,t} \\
&+ \delta_7 * \rho^{2\tau+3} * COV23_{i,t} + \lambda_n * \delta_1' * time + \lambda_n * \delta_2' * \rho^{\tau} * V\Delta EROE1_{i,t} \\
&+ \lambda_n * \delta_3' * \rho^{\tau+1} * V\Delta EROE2_{i,t} + \lambda_n * \delta_4' * \rho^{\tau+2} * V\Delta EROE3_{i,t} \\
&+ \lambda_n * \delta_5' * \rho^{2\tau+1} * COV12_{i,t} + \lambda_n * \delta_6' * \rho^{2\tau+2} * COV13_{i,t} \\
&+ \lambda_n * \delta_7' * \rho^{2\tau+3} * COV23_{i,t} + \xi_{i,t}'
\end{aligned} \tag{14}$$

In (14), λ_n is a dummy variable which is equal to 1 when observations are for newly-listed firms and zero otherwise. Newly-listed firms are defined as firms for which CRSP data has existed for five years or less. All other variables are as previously defined.[27]

Panel A (B) of Table 8 presents results when using (non-)bias-adjusted volatility variables. Although the time-trend for newly listed firms has the expected positive sign in both panes, it is significant for only the 1985-2003 period using one- and two-year horizons of bias-unadjusted forecasts. None of the independent variables is significant for the newly listed firms indicating that the relationship between idiosyncratic volatility and volatility for changes of ROE is consistent between newly listed and more mature firms. Although the increasing trend is higher for newly listed firms, it does not subsume the similar increase in volatility for more mature firms.

[Please place table 8 about here]

A caveat related to our findings is that our sample is inherently more tilted towards large and mature firms for which analysts forecasts are available. However, this bias is reduced in the more recent years as IBES forecasts increase in coverage. Our results are in line with Irvine and Pontiff (2005) who find that mature firms have also experienced increases in volatility. They are also

[26] Equation (14) and subsequently (15) and (16) are estimated using GMM and White standard error correction.

[27] This methodology presents the advantage of not being tainted by sample selection bias which is characteristic of sub-sample investigations.

consistent with Wei and Zhang (2006) who despite their findings that new firms play a higher role in increasing volatility, acknowledge that the total increase in volatility is attributed to both existing and new firms. However, our results differ somewhat from Fink, Fink, Grullon and Weston (2005) who find that firm age is related to lower idiosyncratic risk.

Landsman and Maydew (2002) document an increasing trend in return volatility around earnings announcements. To examine whether the relationship between idiosyncratic volatility and volatility of changes in expected ROE is due to the increase of the volatility characteristics of earnings announcement periods, we estimate the time-series and cross-sectional specification in equation (13) while controlling for earnings announcement dates using a dummy variable.[28] This leads us to the following time-series and cross-sectional regression:

$$\begin{aligned}
IV_{i,t}^{4f} &= \delta_0 + \delta_1 * time + \delta_2 * \rho^{\tau} * V\Delta EROE1_{i,t} + \delta_3 * \rho^{\tau+1} * V\Delta EROE2_{i,t} \\
&+ \delta_4 * \rho^{\tau+2} * V\Delta EROE3_{i,t} + \delta_5 * \rho^{2\tau+1} * COV12_{i,t} + \delta_6 * \rho^{2\tau+2} * COV13_{i,t} \\
&+ \delta_7 * \rho^{2\tau+3} * COV23_{i,t} + \lambda_a * \delta_1'' * time + \lambda_a * \delta_2'' * \rho^{\tau} * V\Delta EROE1_{i,t} \\
&+ \lambda_a * \delta_3'' * \rho^{\tau+1} * V\Delta EROE2_{i,t} + \lambda_a * \delta_4'' * \rho^{\tau+2} * V\Delta EROE3_{i,t} \\
&+ \lambda_a * \delta_5'' * \rho^{2\tau+1} * COV12_{i,t} + \lambda_a * \delta_6'' * \rho^{2\tau+2} * COV13_{i,t} \\
&+ \lambda_a * \delta_7'' * \rho^{2\tau+3} * COV23_{i,t} + \xi_{i,t}''
\end{aligned} \tag{15}$$

In (15), λ_a is a dummy variable which is equal to 1 during the month of the earnings announcement and zero otherwise. All the other variables are as previously defined.

Panels A and B of Table 9 present regression results when using non-bias-adjusted and bias-adjusted forecasts, respectively. In both panels, the time-trend related to earnings announcements is lower and weakly significant compared to the time-trend for periods outside earnings announcements as indicated by the negative and weakly significant coefficients on the dummy time trend. The time-trend coefficient also changes sign but is non-significant when the complete

[28] One avenue for future study is to examine if the relationship is asymmetric when the impact of earnings announcements are conditioned on either the favorableness of the earnings announcements or the surprises contained in the earnings announcements.

forecasts for the one-, two- and three-year horizons are used.[29] The coefficients of the different independent variables are sometimes negative during the announcements periods but are not always significant.

Our results indicate that the upward trend in idiosyncratic volatility and the explanatory power of the independent variables are not exclusive to earnings announcement dates. The significance of the different explanatory variables is robust to earnings announcement dates and seems to be more powerful during non-announcement periods. The lower coefficient values during earnings announcements could be consistent with declining earnings quality and informativeness and/or declining earnings surprises. Furthermore, the increasing trend in idiosyncratic volatility is sometimes even higher during non-announcement periods. This suggests that the increase in volatility might be linked more to the volatility related to fundamental variables for earnings announcement than non-earnings announcement periods. Overall our findings show that the temporal and cross-sectional relationships between idiosyncratic volatility and volatility in changes in expected ROE are robust to earnings announcement dates.

[Please place Table 9 about here]

As a last robustness check, we investigate whether our findings are sensitive to the inclusion of other firm fundamentals that are used extensively in the idiosyncratic volatility literature (namely, size, leverage and BM) using the following regression:

$$\begin{aligned} IV_{i,t}^{4f} &= \delta_0 + \delta_1 * time + \delta_2 * \rho^{\tau} * V\Delta EROE1_{i,t} + \delta_3 * \rho^{\tau+1} * V\Delta EROE2_{i,t} \\ &+ \delta_4 * \rho^{\tau+2} * V\Delta EROE3_{i,t} + \delta_5 * \rho^{2\tau+1} * COV12_{i,t} + \delta_6 * \rho^{2\tau+2} * COV13_{i,t} \\ &+ \delta_7 * \rho^{2\tau+3} * COV23_{i,t} + SZ_{i,t} + LV_{i,t} + BM_{i,t} + \xi_{i,t} \end{aligned} \tag{16}$$

In (16), SZ is measured as the logarithm of market capitalization;[30] LV is measured as the ratio of long-term debt (including debt in current assets) to book value of total assets; and BM is measured as the ratio of book to market value of equity.[31] All the other variables are as previously defined.

[29] Non-reported F-statistics indicate that the total time-trend coefficient is significantly positive during the earnings announcement dates.

As reported in Table 10, size has the expected sign and significance confirming the results of Pastor and Veronesi (2003) that firm size is related to the increasing trend in idiosyncratic volatility. The BM coefficient is not significant whereas the Leverage coefficient is significant but with the opposite sign. Our most important finding is that the relationship between idiosyncratic volatility and volatility in changes in expected ROE is not subsumed by any of these firm fundamentals. This confirms the robustness of the relationship between idiosyncratic volatility and the volatility of the fundamentals.

[Please place Table 10 about here]

8. SUMMARY AND DISCUSSION

Overall, our results indicate that the upward trend in idiosyncratic volatility is associated with a similar increasing trend in the volatility of cash-flow news, as measured by the volatility of monthly changes in expected ROE of one-, two- and three-year horizon forecasts. This relationship is persistent through time and is robust to the inclusion of other fundamental variables (such as size, leverage and book-to-market) and for both newly-listed and mature firms as well as earnings (non-)announcement periods. While investors appear to place greater weights on variables related to forecasts unadjusted for bias, overall model significance improves when correcting for the bias. Investors do not focus only on the information content of the nearest horizon forecast, they also use the two- and three-year forecasts to better assess return volatility. Cash-flow news related to these forecasts generally have their expected sign and are significant both cross-sectionally and across time. However, cross-correlation impacts are not always clearly interpreted by the market and might not always be accounted for when assessing return volatility. Nonetheless, the association between cash-flow news forecasts and idiosyncratic volatility is clearly established in all our findings. Our results are consistent with Wei and Zhang (2006), Irvine and Pontiff (2005), Rajgopal and Venkatachalam (2006) and Jiang and Lee (2006) who use

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[30] Market values are computed at the end of the first quarter following the end of the current fiscal year so that prices reflect all available accounting information for the year.
[31] Book value of equity is measured as total assets minus total liabilities.

different measures of cash flows and/or cash-flow volatility, and find that they are related to idiosyncratic volatility. Our results are also consistent with those of Pastor and Veronesi (2006) who argue that the level and volatility of stock prices are positively linked through firm-specific uncertainty about future profitability.

Our results are consistent with the deteriorating earnings quality and informativeness hypothesis (Lev and Zarowin, 1999; Rajgopal and Venkatachalam, 2006) since a decline in earnings quality and financial statement informativeness leads to an increase in the volatility of cash-flow news. In turn, this is reflected in the increase of idiosyncratic returns. For instance, higher earnings smoothing might mask actual firm profitability, mislead analysts and induce higher volatility in expected earnings news. Similarly, the increasing trend in accounting conservatism (Givoly and Hayn, 2000) could lead to poorer earnings quality when reported earnings fail to provide a good indicator of sustainability. However, since analysts use sources of forecasts other than financial statements, the information content hypothesis could be challenged by the increase in individual risk hypothesis. We examine this latter hypothesis and find that the results do not differ between newly-listed firms (i.e., those with no past earnings, higher risk, and low information quality) and more mature firms (i.e., those with lower risk and an established quality).[32] This suggests that informativeness and the deteriorating earnings quality hypothesis provide more plausible explanations based on our findings.

9. CONCLUSION

Using the Campbell (1991) return decomposition framework, we related idiosyncratic volatility to volatility of changes in expected ROE for one-, two- and three-year horizon forecasts (i.e., the volatility of cash-flow news). We found that the upward trend in idiosyncratic volatility documented by Campbell et al. (2001) is associated with a similar increasing trend in the volatility of cash-flow news for these horizon forecasts. The relationship is not only persistent through time

[32] These results should be interpreted with some caution since our sample is tilted towards more mature firms which results in an unspecified bias against the informativeness hypothesis.

but also robust to the inclusion of other fundamental variables (such as size, leverage and book-to-market) and consistent for both newly-listed and earnings (non-)announcement periods.

Our findings are consistent with the market efficiency hypothesis which suggests that stock return volatility is caused by changes in fundamental variables. Idiosyncratic volatility is related to a similar phenomenon in the volatility of cash-flow news, which persists even after correcting for any bias in analysts' forecasts. Our results favor an informational explanation to the increasing trend in idiosyncratic volatility since a decline in earnings quality and financial statement informativeness leads to an increase in the volatility of cash-flow news. In turn, this is reflected by in increase in the idiosyncratic volatility of stock returns.

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Table 1. Descriptive statistics for the volatilities and covariances of bias-unadjusted changes in EROE (VΔEROE), 1977-2003

This table reports descriptive statistics for the volatilities of monthly changes in expected returns on equity (EROE1, EROE2, and EROE3) for year t, year t+1, and year t+2, respectively. EROE1, EROE2 and EROE3 are measured using monthly IBES EPS forecasts for EPS at the end of year t, t+1 and t+2, respectively, multiplied by the number of common shares, and divided by the value of equity at the end of year t-1. COV12, COV13 and COV23 are the covariances between the changes of forecasts for year t and t+1, t and t+2 and t+1 and t+2, respectively. The sample of forecasts is from 1977 to 2003, except for the 3-year forecasts which start from 1983.

		1977-2005	1976-1980	1981-1985	1986-1990	1991-1995	1996-2000	2001-2003
Panel A: Volatility of equal-weighted and value-weighted unadjusted changes in EROE								
$V\Delta EROE1_{ew}$	Firms	2081	953	1296	1622	2116	3476	2996
	Mean	0.0021	0.0007	0.0014	0.0023	0.0020	0.0030	0.0031
	Stddev	0.0010	0.0002	0.0004	0.0003	0.0002	0.0003	0.0011
	Skew	0.2136	-0.3209	0.7556	0.6022	0.4395	-0.2381	-0.9409
	Kurt	2.4084	2.1208	2.6123	3.5624	2.9522	2.3892	2.7864
$V\Delta EROE2_{ew}$	Firms	1994	884	1244	1546	2041	3335	2883
	Mean	0.0027	0.0004	0.0016	0.0026	0.0031	0.0045	0.0039
	Stddev	0.0015	0.0001	0.0006	0.0004	0.0004	0.0005	0.0015
	Skew	0.0571	0.6002	0.4341	0.3177	0.1193	0.2476	-0.6303
	Kurt	2.0692	2.3993	2.3399	2.4141	2.8118	3.1927	1.9419
$V\Delta EROE3_{ew}$	Firms	932	*NA*	316	359	826	1370	1423
	Mean	0.0035	*NA*	0.0041	0.0023	0.0034	0.0046	0.0034
	Stddev	0.0015	*NA*	0.0025	0.0017	0.0009	0.0007	0.0012
	Skew	-0.1823	*NA*	-0.4015	1.1545	-0.0785	0.6433	-0.6432
	Kurt	2.4977	*NA*	2.2827	3.1499	2.7015	3.2762	2.3119
$V\Delta EROE1_{vw}$	Firms	2081	953	1296	1622	2116	3476	2996
	Mean	0.0008	0.0002	0.0004	0.0007	0.0010	0.0010	0.0011
	Stddev	0.0005	0.0000	0.0002	0.0002	0.0005	0.0003	0.0009
	Skew	1.3047	0.1538	2.1865	1.2874	0.6040	-0.1298	0.2855
	Kurt	4.6209	2.4985	8.0370	5.6125	1.7371	2.4462	2.0443
$V\Delta EROE2_{vw}$	Firms	1994	884	1244	1546	2041	3335	2883
	Mean	0.0012	0.0002	0.0006	0.0010	0.0019	0.0016	0.0016
	Stddev	0.0008	0.0001	0.0002	0.0002	0.0005	0.0004	0.0013
	Skew	0.6776	0.6090	0.3606	1.1279	-0.0844	0.4480	-0.0335
	Kurt	2.8802	3.2487	1.8758	4.2373	1.9205	3.0235	1.4993
$V\Delta EROE3_{vw}$	Firms	932	*NA*	316	359	826	1370	1423
	Mean	0.0015	*NA*	0.0014	0.0010	0.0019	0.0019	0.0013
	Stddev	0.0009	*NA*	0.0009	0.0009	0.0006	0.0006	0.0010
	Skew	-0.0624	*NA*	-0.4298	1.0656	0.7237	1.9045	-0.3287
	Kurt	2.6006	*NA*	2.3075	2.9740	3.2629	6.3404	1.3481

Table 1. Continued.

Panel B: Covariances of unadjusted changes in EROE								
COV12	Mean	0.0004	0.0001	0.0002	0.0003	0.0004	0.0006	0.0009
	Stddev	0.0003	0.0000	0.0001	0.0001	0.0001	0.0001	0.0004
	Skew	1.4460	1.0008	0.5180	0.0791	0.1173	0.1429	-0.4536
	Kurt	4.9819	3.6960	1.7873	2.0957	2.8231	1.8418	2.1657
COV13	Mean	0.0002	*NA*	0.0002	0.0002	0.0002	0.0002	0.0003
	Stddev	0.0001	*NA*	0.0001	0.0001	0.0001	0.0001	0.0002
	Skew	0.7375	*NA*	-0.6495	0.2055	-0.1211	-0.1298	-0.0808
	Kurt	3.6469	*NA*	2.0330	1.9135	2.3285	1.9855	2.1303
COV23	Mean	0.0007	*NA*	0.0012	0.0008	0.0007	0.0008	0.0007
	Stddev	0.0004	*NA*	0.0007	0.0001	0.0003	0.0001	0.0004
	Skew	0.5904	*NA*	-0.8770	0.2607	0.1585	0.3542	-0.3937
	Kurt	3.6963	*NA*	2.0302	2.8061	2.6968	2.9502	2.1074
Panel C: Correlation matrix								
$V\Delta EROE1_{ew}$	1.000							
$V\Delta EROE2_{ew}$	0.801	1.000						
$V\Delta EROE3_{ew}$	0.392	0.531	1.000					
COV12	0.797	0.744	0.335	1.000				
COV13	0.608	0.712	0.581	0.683	1.000			
COV23	0.319	0.363	0.810	0.343	0.592	1.000		

Table 2. Descriptive statistics for the volatilities and covariances of bias-adjusted changes in ROE expectations (V∆AEROE), 1977-2003

This table reports descriptive statistics for the volatility of monthly changes in bias-adjusted expected return on equity (AEROE1, AEROE2, and AEROE3) for year t, year t+1, and year t+2, respectively. EROE is measured using monthly bias-adjusted IBES EPS forecasts for EPS at the end of year t, t+1 and t+2 multiplied by the number of common shares and divided by the value of equity at the end of year t-1. IBES EPS forecasts are corrected for forecast bias using average historical biases as shown in equations (6)-(9). COV12, COV13 and COV23 measure the covariances between the bias-adjusted changes of forecasts for year t and t+1, t and t+2 and t+1 and t+2, respectively. The sample of forecasts is from 1977 to 2003, except for the 3-year forecasts where the bias-adjusted values are computed starting from 1986.

		1977-2005	1976-1980	1981-1985	1986-1990	1991-1995	1996-2000	2001-2003
Panel A: Volatilities of equal- and value-weighted bias-adjusted changes in EROE								
$V\Delta AEROE1_{ew}$	Firms	1331	604	941	1088	1322	1967	2070
	Mean	0.005	0.002	0.004	0.006	0.005	0.006	0.007
	Stddev	0.002	0.001	0.002	0.001	0.000	0.001	0.002
	Skew	-0.277	0.993	-0.365	1.331	0.290	0.180	-1.357
	Kurt	2.726	6.263	1.794	5.071	2.511	1.707	3.816
$V\Delta AEROE2_{ew}$	Firms	1331	604	941	1088	1322	1967	2070
	Mean	0.003	0.001	0.001	0.002	0.003	0.005	0.007
	Stddev	0.002	0.0004	0.001	0.001	0.0004	0.001	0.001
	Skew	0.689	3.808	0.470	1.496	1.152	0.864	-0.201
	Kurt	2.726	17.432	1.831	3.561	3.220	2.757	3.651
$V\Delta AEROE3_{ew}$	Firms	371	*NA*	*NA*	68	310	512	520
	Mean	0.005	*NA*	*NA*	0.003	0.006	0.004	0.005
	Stddev	0.002	*NA*	*NA*	0.002	0.002	0.001	0.002
	Skew	0.658	*NA*	*NA*	-0.147	0.714	0.957	-0.558
	Kurt	3.959	*NA*	*NA*	1.828	1.994	2.811	2.412
$V\Delta AEROE1_{vw}$	Firms	1331	604	941	1088	1322	1967	2070
	Mean	0.0015	0.0004	0.0010	0.0012	0.0031	0.0014	0.0018
	Stddev	0.0014	0.0002	0.0005	0.0002	0.0022	0.0006	0.0012
	Skew	2.2068	2.6674	-0.1068	0.9450	0.6322	1.3480	-0.2405
	Kurt	8.7329	12.9308	1.4657	3.5037	2.1177	4.0802	2.3055
$V\Delta AEROE2_{vw}$	Firms	1331	604	941	1088	1322	1967	2070
	Mean	0.0013	0.0005	0.0006	0.0010	0.0018	0.0017	0.0024
	Stddev	0.0008	0.0004	0.0002	0.0002	0.0005	0.0004	0.0010
	Skew	0.8585	2.6692	0.1417	0.0748	-0.0415	1.0297	-0.4501
	Kurt	3.2359	10.9445	1.8646	2.3053	1.8649	3.7856	2.5432
$V\Delta AEROE3_{vw}$	Firms	371	*NA*	*NA*	68	310	512	520
	Mean	0.0018	*NA*	*NA*	0.0006	0.0030	0.0019	0.0017
	Stddev	0.0014	*NA*	*NA*	0.0005	0.0014	0.0010	0.0009
	Skew	0.8706	*NA*	*NA*	-0.2696	0.1737	1.9431	-0.8308
	Kurt	3.2314	*NA*	*NA*	1.3202	1.4674	6.6664	2.4148

Table 2. Continued.

		1977-2005	1976-1980	1981-1985	1986-1990	1991-1995	1996-2000	2001-2003
Panel B: Covariances of bias-adjusted changes in EROE								
COV12	Mean	0.0003	0.0001	0.0002	0.0003	0.0003	0.0005	0.0006
	Stddev	0.0002	0.0000	0.0001	0.0002	0.0001	0.0001	0.0003
	Skew	1.2523	0.4368	-0.0946	0.0256	-0.6246	0.4060	-0.3700
	Kurt	4.4918	3.0838	2.1960	2.3871	3.9527	1.8196	2.2866
COV13	Mean	0.0002	*NA*	*NA*	0.0002	0.0003	0.0003	0.0003
	Stddev	0.0003	*NA*	*NA*	0.0001	0.0003	0.0001	0.0002
	Skew	0.5915	*NA*	*NA*	3.0317	0.5807	1.5286	-0.4647
	Kurt	3.5098	*NA*	*NA*	21.0601	2.4523	5.2989	2.0812
COV23	Mean	0.0007	*NA*	*NA*	0.0005	0.0008	0.0005	0.0008
	Stddev	0.0003	*NA*	*NA*	0.0002	0.0003	0.0002	0.0004
	Skew	0.1901	*NA*	*NA*	-0.1101	0.7331	0.0982	-0.7270
	Kurt	2.6991	*NA*	*NA*	3.2054	2.9331	2.8181	2.6331
Panel C: Correlation matrix								
$V\Delta AEROE1_{ew}$	1							
$V\Delta AEROE2_{ew}$	0.669	1						
$V\Delta AEROE3_{ew}$	0.148	0.279	1					
COV12	0.718	0.827	0.315	1				
COV13	0.309	0.352	0.577	0.279	1			
COV23	0.413	0.374	0.788	0.428	0.471	1		

Table 3. Descriptive statistics of idiosyncratic volatility measures, 1977-2005

This table reports descriptive statistics for stock return idiosyncratic volatilities as measured using a single market factor model (IV^{1f}) and a 4-factor model (IV^{4f}) based on the Fama and French (1992) 3-factor and the Carhart (1997) 4-factor models. The reported equal- and value-weighted average monthly stock return idiosyncratic volatilities are measured using daily data. Weights are computed using market capitalizations at the end of the previous month. Panel A reports descriptive statistics for monthly return idiosyncratic volatilities for all CRSP firms. Panel B reports descriptive statistics for CRSP firms which satisfy our data requirements and have IBES forecasts.

		1977-2005	1977-1980	1981-1985	1986-1990	1991-1995	1996-2000	2001-2005
Panel A: All firms (N = 1501088)								
	Firms	3714	1566	1467	2465	3989	6593	6346
IV^{1f}_{ew}	Mean	0.022	0.010	0.014	0.024	0.031	0.047	0.033
	Stddev	0.019	0.003	0.004	0.011	0.005	0.020	0.019
	Skew	1.305	1.052	0.339	2.521	0.808	1.113	1.067
	Kurt	5.259	3.576	2.453	9.269	4.616	3.316	3.628
IV^{1f}_{vw}	Mean	0.005	0.003	0.004	0.004	0.006	0.014	0.008
	Stddev	0.006	0.001	0.001	0.003	0.001	0.009	0.005
	Skew	2.929	1.559	0.325	5.230	0.655	1.615	1.635
	Kurt	14.793	6.558	2.375	34.958	3.736	5.298	5.076
IV^{4f}_{ew}	Mean	0.018	0.009	0.012	0.019	0.027	0.040	0.028
	Stddev	0.016	0.002	0.003	0.008	0.004	0.016	0.015
	Skew	1.284	0.947	0.295	2.241	0.701	1.145	1.062
	Kurt	5.243	3.641	2.342	7.462	4.318	3.433	3.562
IV^{4f}_{vw}	Mean	0.004	0.002	0.003	0.003	0.005	0.011	0.006
	Stddev	0.004	0.001	0.001	0.001	0.001	0.007	0.004
	Skew	2.745	0.719	0.259	2.455	0.460	1.404	1.547
	Kurt	13.352	2.828	2.305	10.767	3.321	4.396	4.834
Panel B: Firms with IBES forecasts (N = 674746)								
	Firms	1914	712	767	1226	2259	3535	2971
$IV^{1f}_{ew,ibes}$	Mean	0.021	0.008	0.010	0.017	0.023	0.039	0.031
	Stddev	0.016	0.002	0.002	0.009	0.004	0.019	0.019
	Skew	1.890	0.727	1.075	2.732	1.404	1.109	0.999
	Kurt	6.816	3.068	4.326	10.177	5.982	3.321	3.288
$IV^{1f}_{vw,ibes}$	Mean	0.006	0.003	0.003	0.004	0.005	0.012	0.007
	Stddev	0.005	0.001	0.001	0.003	0.001	0.008	0.005
	Skew	3.157	1.610	0.659	5.335	0.905	1.606	1.668
	Kurt	15.490	7.759	2.983	36.037	4.257	5.389	5.329
$IV^{4f}_{ew,ibes}$	Mean	0.018	0.007	0.008	0.014	0.019	0.032	0.026
	Stddev	0.013	0.002	0.002	0.006	0.004	0.016	0.016
	Skew	1.910	0.543	0.956	2.570	1.351	1.165	1.048
	Kurt	7.038	2.928	4.018	9.375	5.992	3.571	3.394
$IV^{4f}_{vw,ibes}$	Mean	0.004	0.002	0.003	0.003	0.004	0.010	0.006
	Stddev	0.004	0.001	0.001	0.001	0.001	0.006	0.003
	Skew	2.988	0.440	0.633	2.419	0.622	1.363	1.594
	Kurt	13.906	3.865	3.236	10.570	3.583	4.346	5.173

Table 4. Cross-sectional regression results using unadjusted forecasts

This table reports regression results of stock return idiosyncratic volatilities and volatilities of unadjusted monthly changes in forecasts of one-, two- and three-year ROE, as described in equation (12).

	β_0	β_1	β_2	β_3	β_4	β_5	β_6	R^2
Panel A: Results using one- and two-year forecasts								
1977-2003	0.012	0.409	0.427		-0.323			0.05
	(22.82)	(9.46)	(15.84)		(-5.27)			
1977-1980	0.005	0.338	0.836		-1.415			0.05
	(22.85)	(3.63)	(7.43)		(-8.34)			
1981-1985	0.006	0.605	0.419		-0.594			0.07
	(35.30)	(3.34)	(6.34)		(-3.27)			
1986-1990	0.008	0.344	0.264		0.059			0.07
	(13.84)	(6.64)	(5.13)		(0.56)			
1991-1995	0.011	0.460	0.309		-0.204			0.04
	(44.69)	(4.23)	(8.51)		(-1.68)			
1996-2000	0.022	0.334	0.381		0.126			0.03
	(15.72)	(10.92)	(10.25)		(1.77)			
2001-2003	0.022	0.332	0.441		0.004			0.03
	(11.97)	(7.48)	(9.89)		(0.04)			
Panel B: Results using one-, two- and three-year forecasts								
1985-2003	0.011	0.221	0.439	0.167	-0.307	-0.017	-0.463	0.11
	(17.06)	(4.79)	(13.94)	(6.60)	(-2.77)	(-0.10)	(-3.37)	
1986-1990	0.004	0.176	0.474	0.024	-0.817	0.138	-0.814	0.11
	(19.39)	(1.39)	(7.48)	(0.38)	(-2.85)	(0.27)	(-2.01)	
1991-1995	0.006	0.194	0.278	0.220	0.160	-0.131	-0.278	0.11
	(32.84)	(4.63)	(7.22)	(6.85)	(1.35)	(-1.24)	(-3.19)	
1996-2000	0.019	0.301	0.469	0.207	-0.128	0.053	-0.282	0.09
	(14.24)	(4.78)	(8.66)	(6.96)	(-0.91)	(0.26)	(-2.15)	
2001-2003	0.019	0.224	0.582	0.301	-0.334	-0.256	-0.356	0.05
	(11.50)	(3.51)	(5.89)	(5.73)	(-1.66)	(-2.07)	(-3.24)	

Table 5. Cross-sectional regression results using bias-adjusted forecasts

This table reports regressions of stock return idiosyncratic volatilities and volatilities of bias-adjusted monthly changes in forecasts of one-, two- and three-year ROE, as described in equation (12). The bias in the forecasts of analysts is adjusted using the model developed in equations (6) to (9).

	β_0	β_1	β_2	β_3	β_4	β_5	β_6	R^2
Panel A: Results using one-year and two-year forecasts								
1977-2003	0.009	0.145	0.241		-0.224			0.05
	(24.28)	(5.80)	(8.43)		(-5.07)			
1977-1980	0.004	0.117	0.381		-0.658			0.04
	(27.18)	(5.32)	(2.70)		(-3.21)			
1981-1985	0.005	0.225	0.289		-0.434			0.09
	(28.62)	(1.72)	(2.97)		(-4.51)			
1986-1990	0.006	0.097	0.210		-0.145			0.04
	(16.76)	(6.79)	(5.45)		(-1.73)			
1991-1995	0.008	0.121	0.164		-0.099			0.04
	(40.81)	(6.67)	(7.95)		(-1.24)			
1996-2000	0.015	0.158	0.198		0.052			0.05
	(15.10)	(12.57)	(13.05)		(0.84)			
2001-2003	0.016	0.148	0.228		-0.123			0.04
	(13.76)	(10.90)	(8.18)		(-2.45)			
Panel B: Results using one-, two- and three-year forecasts								
1985-2003	0.007	0.174	0.290	0.059	0.247	-1.367	0.507	
	(15.79)	(3.42)	(7.86)	(1.89)	(0.96)	(-0.99)	(0.47)	0.14
1986-1990	0.003	0.522	0.351	-0.087	1.275	-1.336	0.864	0.17
	(9.446)	(2.142)	(2.398)	(-0.595)	(1.045)	(-0.892)	(0.516)	
1991-1995	0.004	0.073	0.195	0.044	0.283	-0.139	-0.056	0.15
	(20.84)	(4.09)	(4.99)	(3.21)	(1.56)	(-2.28)	(-1.26)	
1996-2000	0.010	0.127	0.279	0.149	0.009	-0.177	-0.199	0.13
	(12.15)	(4.50)	(9.59)	(8.36)	(0.09)	(-2.25)	(-2.64)	
2001-2003	0.013	0.051	0.403	0.089	-0.456	-0.200	0.148	0.09
	(10.99)	(0.94)	(4.36)	(1.69)	(-1.22)	(-1.24)	(1.42)	

Table 6: Time-Series regression results using bias (un)adjusted forecasts

This table reports time-series and cross-section results of stock return idiosyncratic volatilities and volatilities of monthly changes in forecasts of one-, two- and three-year ROE, as described in equation (13).

	Time*10^5	δ_2	δ_3	δ_4	δ_5	δ_6	δ_7	R^2
Panel A: Results using bias-unadjusted forecasts								
1977-2003	3.170	0.109	0.069		0.04			0.20
	(16.05)	(3.33)	(3.21)		(0.59)			
1985-2003	5.080	0.110	0.069		0.04			0.20
	(15.90)	(3.23)	(3.17)		(0.52)			
1985-2003	3.780	0.099	0.094	0.026	0.092	0.086	-0.061	0.21
	(11.28)	(1.71)	(3.29)	(0.96)	(0.98)	(0.58)	(-0.77)	
Panel B: Results using bias-adjusted forecasts								
1977-2003	2.300	0.046	0.061		-0.081			0.25
	(13.54)	(4.43)	(4.04)		(-1.71)			
1985-2003	3.590	0.045	0.060		-0.085			0.25
	(13.61)	(4.27)	(3.82)		(-1.74)			
1985-2003	4.850	0.029	0.093	0.046	-0.021	-0.032	-0.099	0.28
	(8.81)	(1.74)	(3.06)	(1.05)	(-0.31)	(-0.49)	(-1.40)	

Table 7: Panel unit root tests

This table presents unit root test results for the volatilities of monthly changes in forecasts of one-, two- and three-year ROE forecasts for bias (un)adjusted variables. All observations are structured into panels to better capture the impact of persistence in individual series. The three tests are Levin, Lin and Chu (LLC); Breitung; and Im, Pesaran, and Shin (IPS).

	$IV^{4f}_{ew,ibes}$	$V\Delta EROE_1$	$V\Delta EROE_2$	$V\Delta EROE_3$	$V\Delta AEROE_1$	$V\Delta AEROE_2$	$V\Delta AEROE_3$
LLC	-177.50 (0.00)	-95.42 (0.00)	-120.49 (0.00)	-2.61 (0.00)	108.44 (1.00)	-113.20 (0.00)	-171.52 (0.00)
Breitung	-62.56 (0.00)	-6.34 (0.00)	-7.25 (0.00)	-0.27 (0.39)	-5.98 (0.00)	-31.48 (0.00)	-2.59 (0.00)
IPS	-365.90 (0.00)	-55.85 (0.00)	-96.84 (0.00)	-35.66 (0.00)	-0.48 (0.31)	-11.35 (0.00)	-35.52 (0.00)

Table 8: Time-series and cross-section regression results controlling for newly listed firms

This table reports time-series and cross-section results of stock return idiosyncratic volatilities and volatilities of monthly changes in forecasts of one-, two- and three-year ROE, as described in equation (14).

	Time $*10^5$	δ_2	δ_3	δ_4	δ_5	δ_6	δ_7	λ_n *Time $*10^5$	$\lambda_n * \delta_2'$	$\lambda_n * \delta_3'$	$\lambda_n * \delta_4'$	$\lambda_n * \delta_5'$	$\lambda_n * \delta_6'$	$\lambda_n * \delta_7'$	R^2
Panel A: Results using bias-unadjusted forecasts															
1977-	3.120	0.103	0.073		0.089			0.249	0.051	-0.015		-0.206			0.23
2003	(15.58)	(2.82)	(2.85)		(1.12)			(1.56)	(0.72)	(-0.41)		(-1.64)			
1985-	5.03	0.104	0.074		0.087			0.268	0.051	-0.014		-0.212			0.21
2003	(15.66)	(2.72)	(2.85)		(1.08)			(1.64)	(0.71)	(-0.38)		(-1.68)			
1985-	3.74	0.056	0.097	0.026	0.146	0.084	-0.050	0.199	-0.097	0.029	0.005	-0.106	-0.089	-0.049	0.20
2003	(11.00)	(1.97)	(3.09)	(0.73)	(1.51)	(0.45)	(-0.50)	(1.00)	(-0.84)	(0.50)	(0.09)	(-0.77)	(-0.45)	(-0.30)	
Panel B: Results using bias-adjusted forecasts															
1977-	2.26	0.039	0.072		-0.064			0.151	0.028	-0.034		-0.060			0.24
2003	(13.16)	(3.98)	(3.97)		(-1.09)			(1.06)	(0.95)	(-1.48)		(0.75)			
1985-	2.64	0.041	0.071		-0.066			0.179	0.030	-0.035		-0.062			0.24
2003	(13.45)	(3.83)	(3.88)		(-1.10)			(1.24)	(0.96)	(-1.42)		(-0.76)			
1985-	4.780	0.033	0.103	0.041	0.028	-0.091	-0.059	0.132	-0.011	-0.083	0.028	0.001	0.228	-0.193	0.28
2003	(8.78)	(1.34)	(3.05)	(0.87)	(0.36)	(-1.19)	(-0.836)	(0.156)	(-0.24)	(-1.24)	(0.56)	(0.05)	(1.97)	(-1.66)	

Table 9: Time-series and cross-section regression results controlling for earnings announcements dates

This table reports time-series and cross–section regression results of stock return idiosyncratic volatilities and volatilities of monthly changes in forecasts of one-, two- and three-year ROE, as described in equation (15).

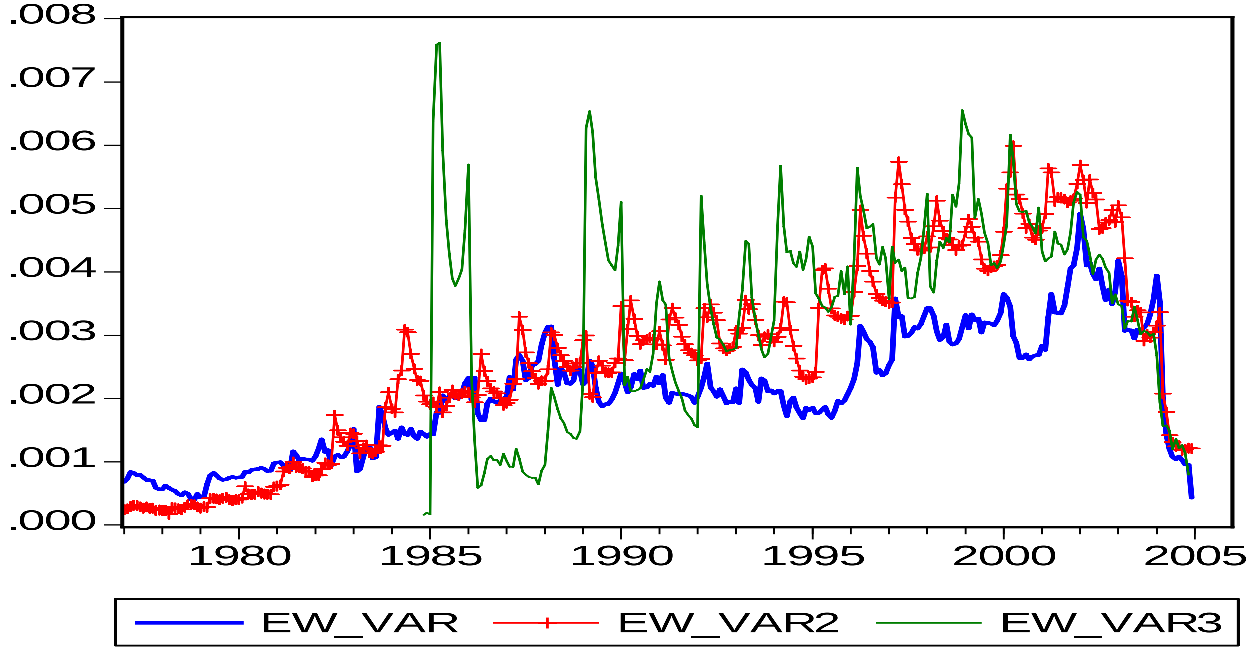
	Time*10^5	δ_2	δ_3	δ_4	δ_5	δ_6	δ_7	λ_a *Time*10^5	$\lambda_a * \delta_2'$	$\lambda_a * \delta_3'$	$\lambda_a * \delta_4'$	$\lambda_a * \delta_5'$	$\lambda_a * \delta_6'$	$\lambda_a * \delta_7'$	R^2
Panel A: Results using bias-unadjusted forecasts															
1977-2003	3.200	0.107	0.073		0.038			-0.217	0.018	-0.042		0.063			0.21
	(16.04)	(3.13)	(3.36)		(0.50)			(-1.66)	(0.34)	(-1.19)		(0.65)			
1985-2003	5.110	0.108	0.074		0.034			-0.212	0.018	-0.043		0.058			0.20
	(15.90)	(3.03)	(3.31)		(0.44)			(-1.61)	(0.33)	(-1.21)		(0.60)			
1985-2003	3.790	0.030	0.100	0.032	0.084	0.096	-0.07	0.927	-0.018	-0.040	-0.082	-0.080	-0.109	0.158	0.20
	(11.21)	(1.71)	(3.23)	(1.15)	(0.85)	(0.62)	(-0.86)	(0.76)	(-0.311)	(-1.26)	(-2.15)	(0.73)	(-0.63)	(1.49)	
Panel B: Results using bias-adjusted forecasts															
1977-2003	2.320	0.041	0.063		-0.077			-0.138	0.017	-0.015		-0.039			0.26
	(13.64)	(4.17)	(4.06)		(-1.54)			(-1.60)	(0.95)	(-0.64)		(-0.41)			
1985-2003	3.61	0.042	0.063		-0.081			-0.141	0.019	-0.016		-0.043			0.25
	(13.67)	(4.02)	(3.85)		(-1.57)			(-1.61)	(1.01)	(-0.67)		(-0.44)			
1985-2003	4.81	0.032	0.098	0.056	-0.043	-0.040	-0.106	0.213	-0.018	-0.044	-0.116	0.148	0.067	0.145	0.29
	(8.87)	(1.77)	(3.06)	(1.26)	(-0.60)	(-0.61)	(-1.44)	(1.49)	(-0.82)	(-2.70)	(-3.16)	(1.67)*	(0.56)	(1.41)	

Table 10: Time-series and cross-section regression results controlling for Size, Leverage, and BM

This table reports time-series and cross-section regression results of stock return idiosyncratic volatilities and volatilities of monthly changes in forecasts of one-, two- and three-year ROE, as described in equation (16).

	Time *10^5	δ_2	δ_3	δ_4	δ_5	δ_6	δ_7	size*10	Lev	BM	R^2
Panel A: Results using bias-unadjusted forecasts											
1977-	3.02	0.187	0.250		-0.05			0.007	0.001	0.001	0.12
2003	(18.08)	(8.08)	(11.41)		(-0.81)			(2.55)	(2.33)	(-1.53)	
1985-	8.860	0.076	0.124		-0.06			-0.007	-0.003	0.004	0.13
2003	(9.74)	(2.10)	(4.07)		(-0.81)			(-5.99)	(-3.26)	(-1.06)	
1985-	3.27	0.190	0.277	0.069	0.137	-0.196	-0.211	-0.003	-0.002	-0.001	0.15
2003	(17.38)	(6.04)	(9.44)	(8.60)	(1.56)	(-2.67)	(-4.92)	(-8.80)	(-6.09)	(-1.13)	
Panel B: Results using bias-adjusted forecasts											
1977-	1.87	0.070	0.115		0.018			-0.002	-0.002	-0.001	0.12
2003	(17.45)	(6.22)	(6.45)		(0.32)			(-6.36)	(-3.43)	(-1.17)	
1985-	2.68	0.080	0.124		0.057			-0.002	-0.002	-0.001	0.10
2003	(20.81)	(7.27)	(7.00)		(1.06)			(-6.45)	(-5.10)	(-1.56)	
1985-	3.18	0.050	0.077	0.027	0.117	-0.074	0.001	-0.001	-0.002	-0.001	0.12
2003	(15.44)	(4.31)	(3.95)	(2.30)	(2.81)	(-2.89)	(0.05)	(-3.69)	(-5.33)	(-1.56)	

Panel A: Equal-weighted volatilities of monthly changes in unadjusted forecasts of ROE



Panel B: Value-weighted volatilities of monthly changes in unadjusted forecasts of ROE

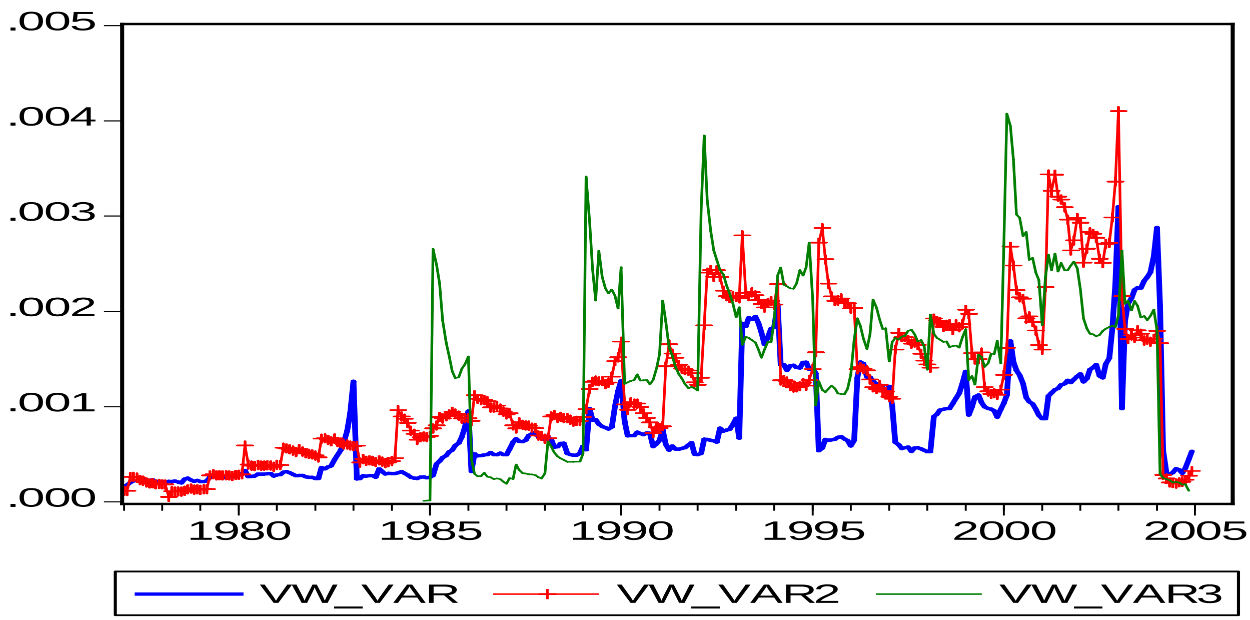
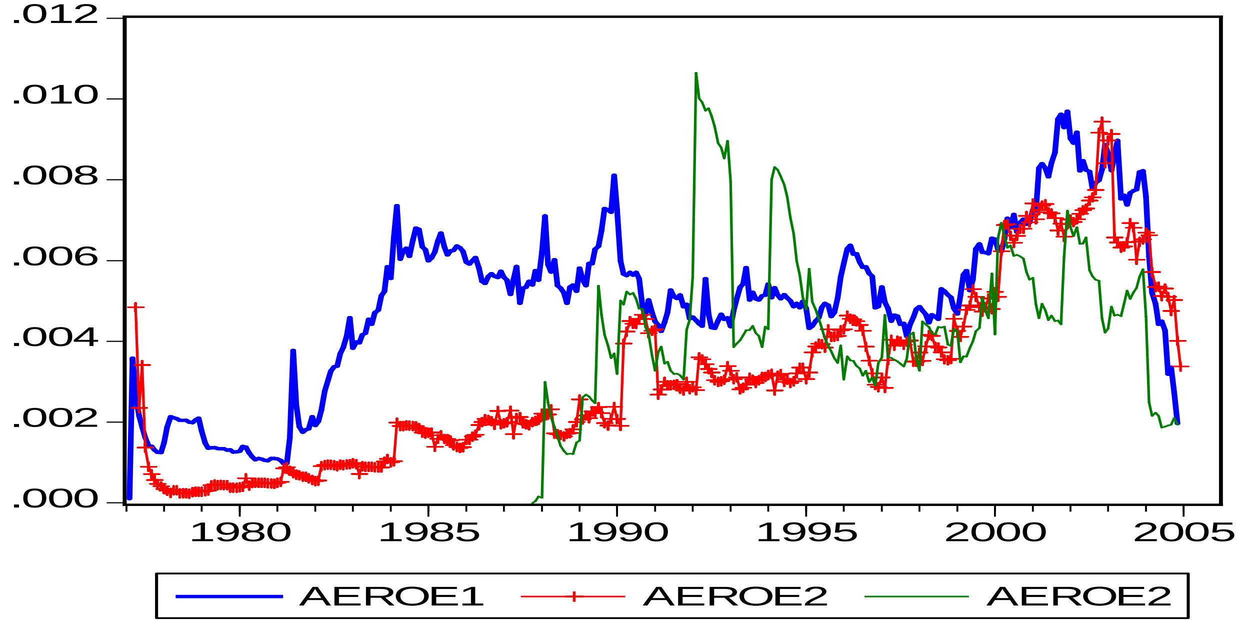


Figure 1: Volatility of monthly changes in unadjusted forecasts of ROE

Panel A: Equal-weighted volatilities of monthly changes in bias-adjusted forecasts of ROE



Panel B: Value-weighted volatilities of monthly changes in bias-adjusted forecasts of ROE

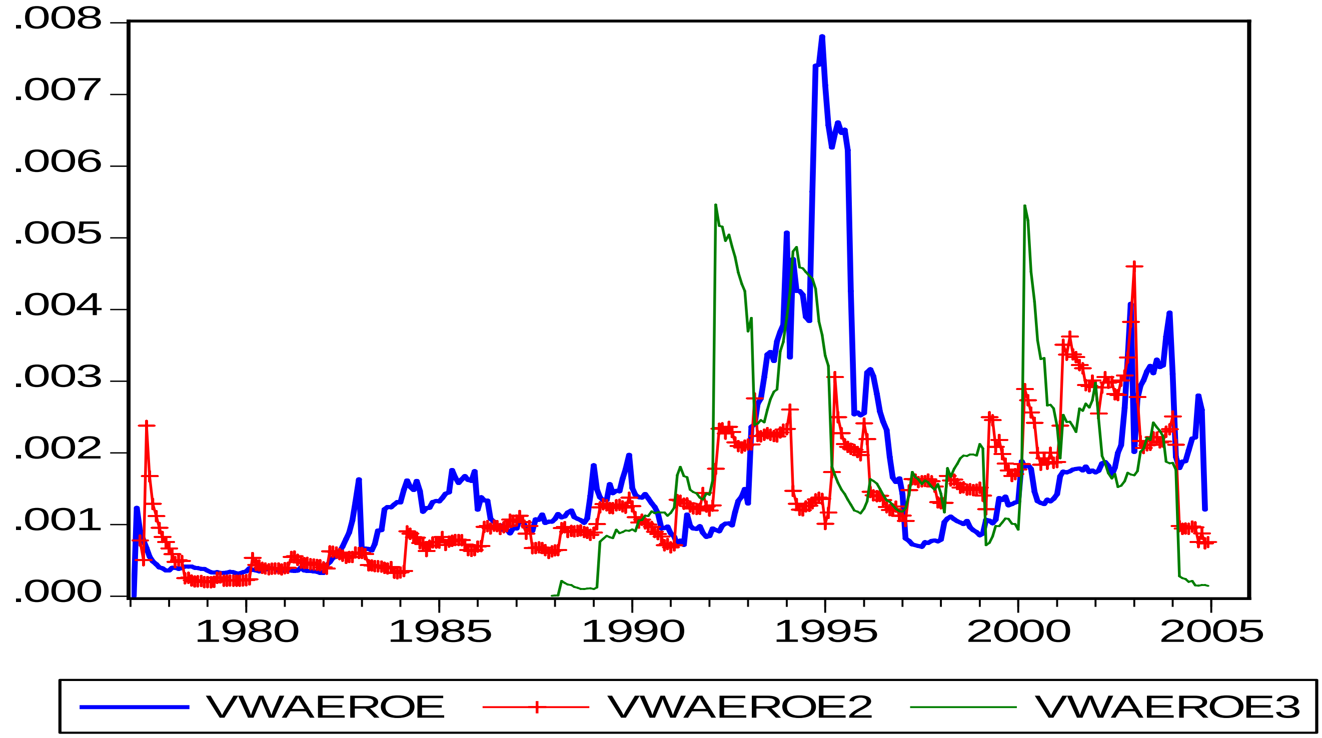
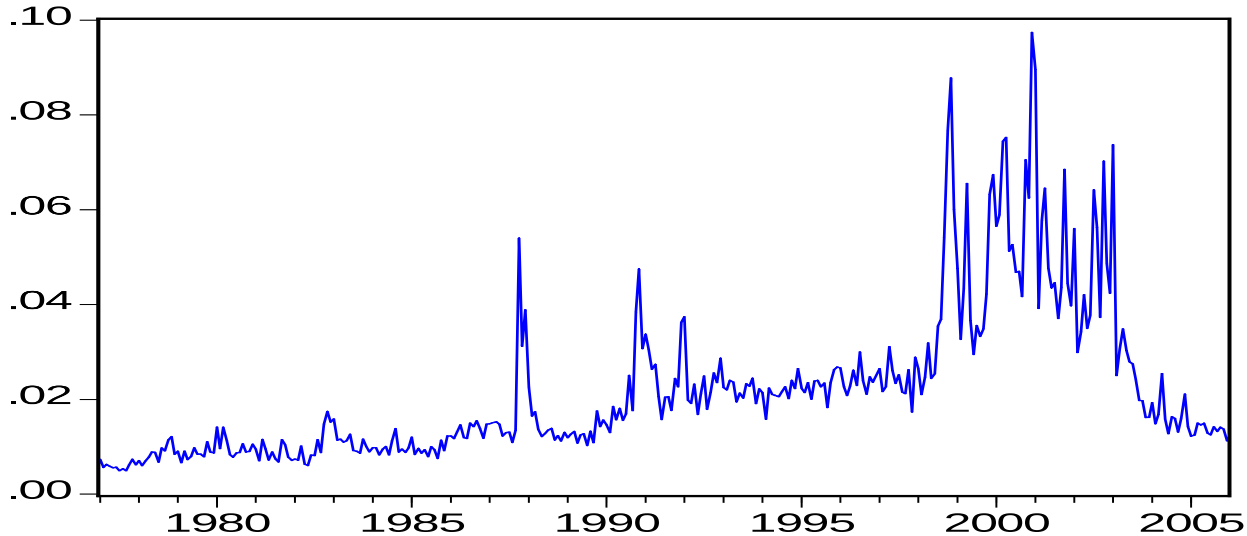


Figure 2: Volatility of monthly changes in bias-adjusted forecasts of ROE

Equal-weighted monthly idiosyncratic volatilities, $IV^{1f}_{ew,ibes}$



Value-weighted monthly idiosyncratic volatilities, $IV^{1f}_{vw,ibes}$

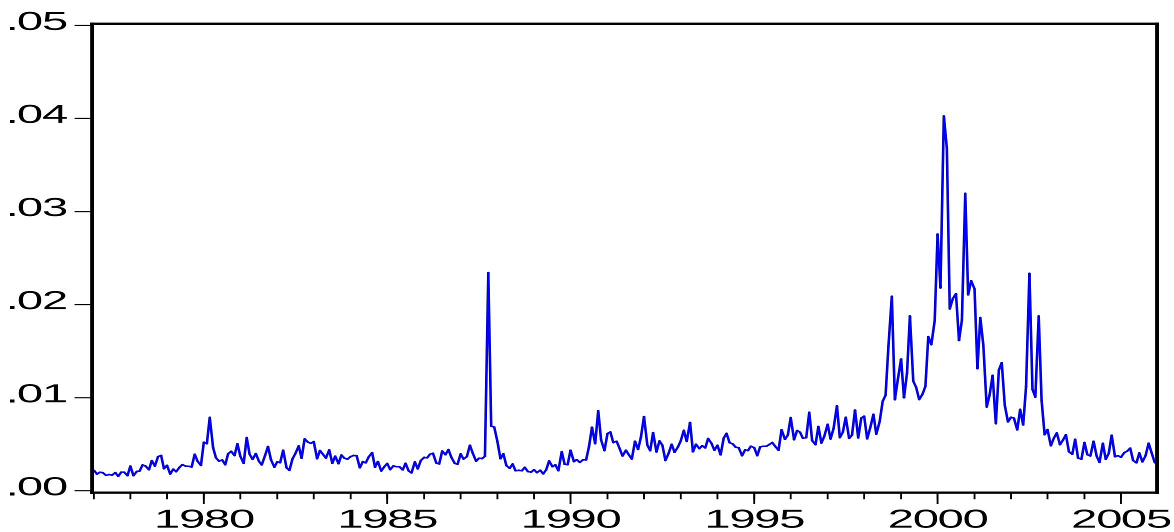


Figure 3: Average monthly return idiosyncratic volatilities